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Explaining International Variations in Self-Employment: Evidence from a Panel of OECD Countries

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Organisation for Economic Cooperation and Development (OECD) data from 1972 to 1996 reveals substantial differences in the levels and trends of self-employment rates across countries. This article uses recently developed panel integration and cointegration techniques to explore the determinants of aggregate self-employment rates. We find that within the panel, self-employment rates are positively and significantly related to personal income tax rates and negatively and significantly related to the unemployment benefit replacement rate. This accords a central role to government tax and transfer policies, in contrast to nonrobust influences from macroeconomic variables, which have been widely used in previous studies.

JEL Classification: C23, J23, H24, H25

1. Introduction

In recent years there has been growing awareness of the importance of self-employment for growth and employment creation. Governments around the world are increasingly implementing policies designed to promote self-employment (see, e.g., OECD 1998). Yet relatively little is known about the determinants of self-employment, especially the effects on the self-employment rate of government policy instruments. As we show in this article, self-employment rates in the OECD display marked variations across countries, both in cross-section "snapshots" and over time. The objective of this article is to explain these disparate patterns by identifying the determinants of self-employment rates, placing special emphasis on government tax and transfer policies.

Previous studies of the determinants of national self-employment rates have been confined to a handful of countries.[1] Although they have been able to shed some light on the causes of self-employment rates within particular countries, they suffer from two drawbacks. First, they cannot explain the pronounced observed *differences* in self-employment rates between countries. Second, national time-series studies tend to work with only short spans of data, consigning tests of statistical significance to lack of power, and hence reliability.

Both of these problems can be addressed by exploiting the panel nature of available OECD data.

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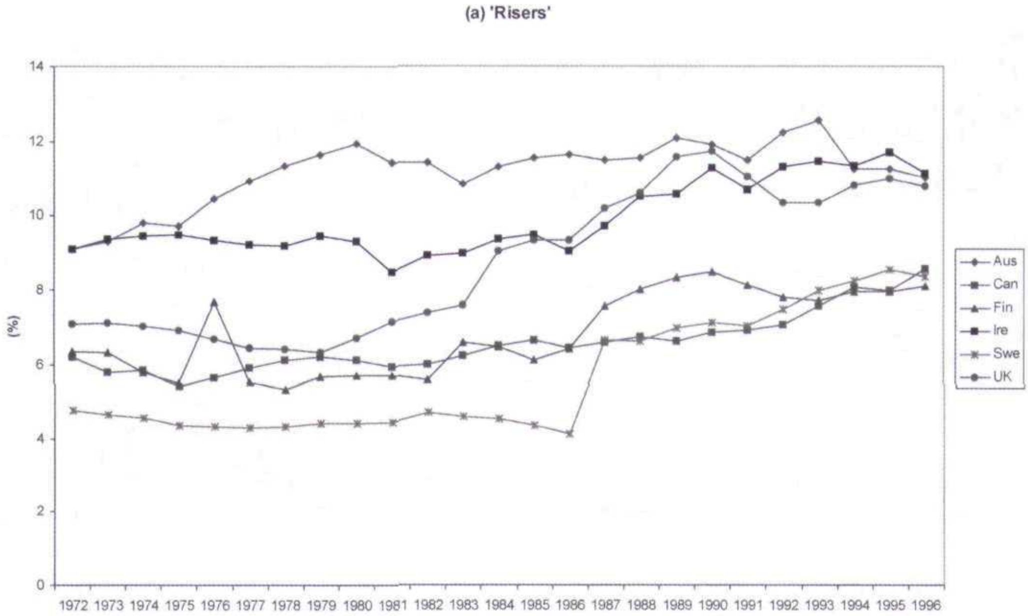
[1] See, e.g., Blau (1987), Evans and Leighton (1989), Alba-Ramirez (1994), and Devine (1994) for U.S. studies; and Rees and Shah (1986); Robson (1991, 1998); Parker (1996); and Cowling and Mitchell (1997) for U.K. studies. Research on other countries includes Foti and Vivarelli (1994) for Italy, Alba-Ramirez (1994) for Spain, and Bernhardt (1994) for Canada.



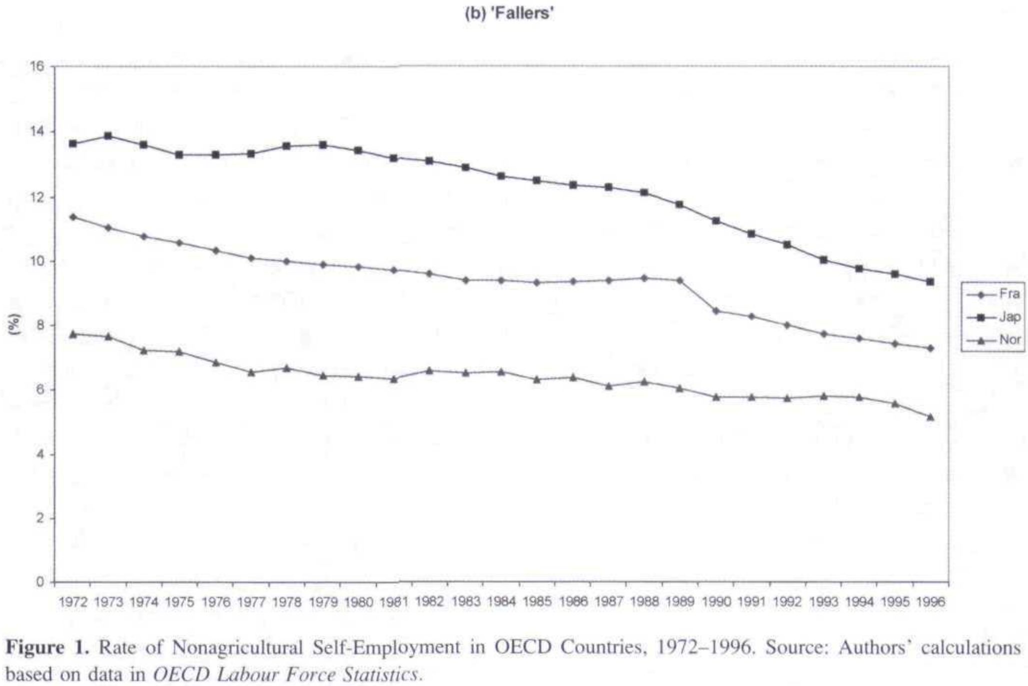


Figure 1. Rate of Nonagricultural Self-Employment in OECD Countries, 1972–1996. Source: Authors' calculations based on data in *OECD Labour Force Statistics*.

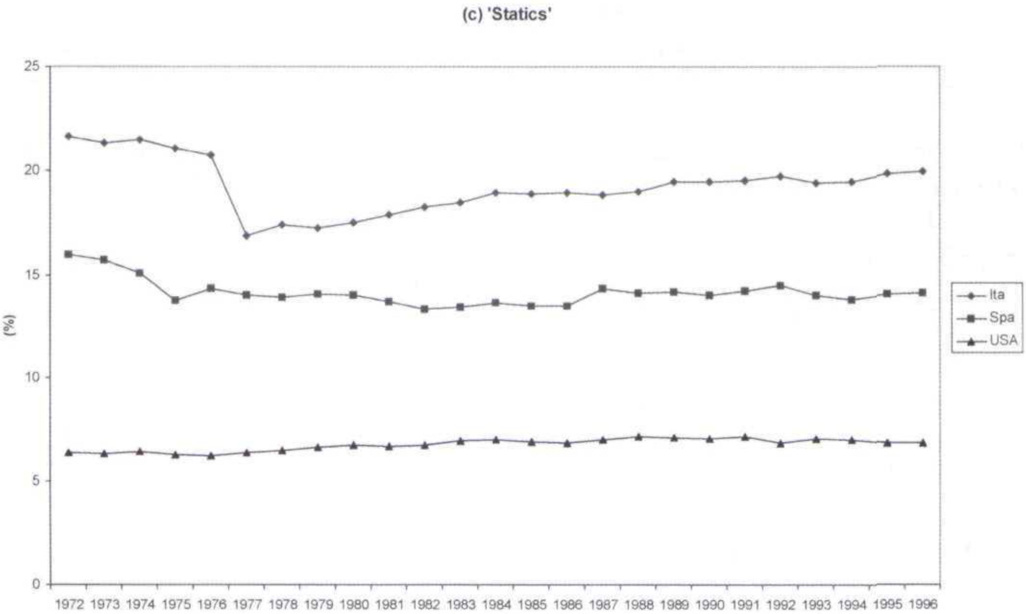


Figure 1. Continued

Panel data enjoys the advantage over static cross-sections or single-country time-series data of looking at more than just one time period and country. Trends in self-employment rates and cross-country differences in these rates are both of interest in their own right. This motivated the influential article of Acs, Audretsch, and Evans (1994; henceforth AAE), as well as Staber and Bogenhold (1993), Robson and Wren (1999), Blanchflower (2000), and OECD (2000). These articles all used ordinary least squares (OLS) to estimate self-employment regressions based on pooled cross-section time-series or fixed/random effects specifications. Yet recent developments in the analysis of panel data regression models cast doubt on the validity of the findings from these studies. Although it has long been recognized that OLS yields biased and inconsistent estimates in dynamic panel data regression analysis (see Nickell 1981), recent work has shown that OLS will also produce biased and inconsistent estimates even in regular panel data models when—as is shown to be the case here—variables possess unit roots.[2] In addition, conventional significance tests based on OLS estimates cannot be used to reliably identify genuine relationships between variables. This problem is well known in the traditional time-series econometrics literature (e.g., Phillips 1986), where it prompted the development of cointegration estimators (Engle and Granger 1987; Johansen 1988). Only recently have these econometric techniques been extended to the panel framework. A key advantage of these techniques is that by utilizing cross-country information, panel unit root and cointegration tests are much more powerful than for the conventional single-country case, making inference more reliable. This point is especially important in view of the low power of conventional unit root and cointegration tests (Banerjee et al. 1993).[3]

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[2] See Pesaran and Smith (1995), Kao (1999), Harris and Tzavalis (1999), and Pedroni (1999b) for formal demonstrations of this point. Moreover, incorporating deterministic time dummies in the panel regressions cannot circumvent the inconsistency of OLS.

[3] The practical importance of using a panel data cointegration estimator in the context of international self-employment rates is illustrated by a recent study by the OECD (2000). When several pooled self-employment regressions were estimated by OLS, virtually no explanatory variable was found to be significant, despite strong evidence to the contrary from national studies. But panel data cointegration techniques conducted for the OECD by one of the authors (S.C.P.) revealed that there were, in fact, significant relationships between the variables (see OECD 2000).

In this article we investigate the determinants of self-employment using a panel of annual data on 12 OECD countries spanning the period 1972–1996. We use a wider range of explanatory variables than previous studies, paying particular attention to variables under the direct control of governments: average rates of personal income tax, employers' social security contributions, and benefit replacement rates. We find that the emphasis on macroeconomic and demographic variables in previous studies appears to have been misplaced. Macroeconomic variables are found to be neither significant nor robust determinants of self-employment rates in the OECD. Instead, government policy variables appear to play a central role. In particular, we show that self-employment rates are positively and significantly related to average income tax rates and negatively and significantly related to the benefit replacement rate. We also show that panel OLS would have failed to uncover these results.

The article is organized as follows. Section 2 documents the variation in national self-employment rates in the panel, describes the data, and presents several hypotheses that may explain these patterns. Section 3 briefly describes the panel unit root and cointegration techniques and presents the results. Section 4 concludes the article.

2. Data and Possible Explanations

Annual data for the rate of nonagricultural self-employment in 12 OECD countries over the period 1972–1996 are plotted in Figure 1 using various issues of *OECD Labor Force Statistics*. The numerator is the number of employers and own account workers in nonagricultural civilian employment, whereas the denominator includes all persons in civilian employment in the nonagricultural sector plus the numbers in unemployment. The agricultural sector is excluded, as self-employment rates in this sector are likely to be heavily influenced by historically and culturally determined traditions of family ownership and factors other than those that influence self-employment rates in the rest of the economy.[4]

The graphs show considerable dispersion in the rate of nonagricultural self-employment in the OECD, ranging from a low of just over 4% for much of the period in Sweden, to an average of around 19% in Italy. There appear to be three distinct groupings of countries. One group, graphed in of Figure 2, experienced a trend increase in the rate of self-employment over the period (Australia, Canada, Finland, Ireland, Sweden, and the United Kingdom). A second group (Figure 1b) experienced a declining rate of self-employment (France, Japan, Norway), and in the third group (Figure 1c), the rate of self-employment remained fairly static (Italy, Spain, and the United States). The sharpest increases occurred in Sweden and the United Kingdom, whereas the steepest declines in the self-employment rate were experienced in Japan and France.

This picture of rather disparate trends and patterns in OECD self-employment echoes that found by AAE in their study for the period 1966–1987 and in the more recent study by Blanchflower (2000).[5] What kind of factors can we identify to try to explain the cross-national variations in self-employment that we observe? A number of potential explanatory variables are suggested by the

[4] Our definition of self-employment excludes unpaid family workers. These comprise a relatively high proportion of self-employment in agriculture (Blanchflower 2000), but are much less commonly observed outside the agricultural sector.

[5] The study by Blanchflower (2000) shows the importance of the inclusion, or otherwise, of the agricultural sector to the identification of trends in the rate of self-employment. When the agricultural sector is included, the trend in the rate of self-employment in OECD countries over the period 1966–1996 is almost uniformly negative, with only the United Kingdom, Portugal, and New Zealand (the latter two are not included in our data set) experiencing any increase. The implication is that within the OECD countries, the rate of self-employment has been falling most substantially in the agricultural sector. As indicated above, the reasons for this decline are likely to be peculiar to this sector and lie outside the scope of the present study.

Table 1. Self-Employment Rates[a] by Gender, 1990 and 1996

	1990		1996	
	Men	Women	Men	Women
United States	10.4	6.2	9.7	6.7
Canada	17.6	9.2	19.8	11.8
Japan	16.4	10.9	14.1	8.5
Australia	16.7	10.9	16.9	10.7
Finland	17.4	10.2	17.7	9.4
France	—	—	14.3[b]	6.1[b]
Ireland	29.9	7.6	27.0	8.1
Italy	—	—	27.8	15.5
Norway	—	—	10.5	4.3
Spain	22.0	15.2	23.0	15.8
Sweden	12.7	4.7	15.3	5.5
United Kingdom	17.7	7.4	17.0	7.0

Sources: International Labor Office, *Yearbook of Labour Statistics* 2000; Eurostat, *Labour Force Survey* and *Labour Force Survey, Historical Supplement*, ONS.

[a] Employers plus own-account workers as a percentage of all in employment.

[b] 1997.

previous literature on this issue. For example, the findings of AAE suggested that the self-employment rate is related to the level of real per capita GDP, the demographic composition of the labor force, and the sectoral composition of GDP. Higher per capita GDP might be related negatively to aggregate self-employment rates if it is associated with greater capital per worker, and hence greater average firm size (Lucas 1978). On the other hand, higher per capita GDP might indicate buoyant demand conditions within countries, which might disproportionately benefit the self-employed. It is therefore not possible to unambiguously sign the effect of per capita GDP on self-employment rates *a priori*.

AAE reported a negative relationship between the self-employment rate and the rate of female labor-force participation. This is consistent with the evidence that self-employment rates tend to be lower among women than men (see Table 1).[6] We would expect a similar relationship to apply in our data. AAE also reported a positive relationship between the self-employment rate and the service sector share of GDP. This may be explained by technological factors that give the self-employed a comparative advantage in the service sector. This is evident from the figures presented in Table 2, which show self-employment rates by sector for selected countries in our data sample. Thus, we predict a positive effect from the service sector share of GDP on self-employment rates.

A number of studies suggest that the rate of self-employment may be related to the rate of unemployment. Two contrasting effects may be at work in this relationship. On the one hand, individuals may be pushed into self-employment by a shortage of opportunities for paid work ("recession push"). In this case, we would expect to see a positive relationship between the rate of unemployment and the rate of self-employment. On the other hand, a high rate of unemployment may be associated with relatively low levels of demand for the output of the self-employed ("prosperity pull"), so that a negative relationship may be observed between these two variables. Individuals may

[6] Some studies have shown significant differences between men and women in the effect of personal characteristics on the likelihood of self-employment (e.g., Evans and Leighton 1989; Burke, Fitzroy, and Nolan 2002). However, the limited evidence that exists suggests that, for the United States at least, the effects of macroeconomic factors on male and female self-employment are broadly similar (Evans and Leighton 1989). Hence, it seems reasonable to aggregate over the two gender groups for the purposes of our analysis.

Table 2. Self-Employment Rates[a] by Sector, Selected Countries, 1997

	Industry	Services
United States[bc]	8.6	10.0
Canada[c]	12.9	17.4
Finland	9.6	10.4
France	8.6	9.1
Ireland	11.7	13.6
Italy	16.6	25.8
Spain	15.4	19.6
Sweden	10.1	9.4
United Kingdom	14.0	11.0

Sources: EU countries, Eurostat, *Labour Force Survey*; United States and Canada, authors' calculations based on data in Manser and Picot (1999).

[a] Self-employment as a percentage of total employment.

[b] 1996.

[c] Includes Transport and Communications.

also feel more comfortable taking on the risks associated with self-employment against the backdrop of a buoyant labor market that offers them the chance of a reasonably quick return to paid employment in the event of business failure. This again would lead us to expect to see a negative relationship between unemployment and self-employment.

Evidence from cross-country studies on this issue is mixed. Staber and Bogenhold (1993) find a positive relationship between the unemployment rate and the rate of self-employment in 17 OECD countries. Blanchflower (2000), however, reports a negative relationship for most of the countries in his data sample. AAE report a positive relationship between the rate of unemployment and the rate of self-employment in a bivariate context, but this disappears when additional regressors are introduced into the equation. Yet it could be argued that if the researcher controls for levels of demand (and thereby the prosperity pull effect), then only the positive recession push effect will be identified. Since we control for aggregate income in all of our estimations, we tentatively predict a positive effect from unemployment rates on self-employment rates.

As well as using the variables outlined above, we also consider the effect on cross-national variations in self-employment of three tax and benefit variables. These are income tax and employees' social security contributions as a percentage of personal income (the average rate of income tax); employers' contributions to social security as a percentage of wages and salaries (the rate of payroll tax); and an OECD summary measure of the ratio of unemployment benefits to earnings (the replacement rate).

The possible effects of income tax rates on participation in self-employment have been well documented in the literature: see, for example, Blau (1987); Parker (1996, 1999, 2001, 2003); Robson and Wren (1999); Bruce (2000); and Scheutze (2000). High rates of income tax may in principle have both positive and negative effects on the incentive for self-employment. The greater opportunities that are generally available to self-employed workers (relative to wage and salary workers) both for tax deduction of work-related expenses and for income tax evasion tend to favor a positive relationship between tax rates and self-employment. However, the tendency for high tax rates to diminish the incentive to supply effort may reduce the incentive for self-employment. In general, most of the empirical literature tends to find that the former effect dominates the latter, implying that higher tax rates are generally found to lead to an increase in self-employment. Part of the reason for the dominance of positive effects might be that marginal income tax rates are the same or very similar for employees and the self-employed in most countries (Price Waterhouse 2002). Therefore, labor supply effects of changing tax rates are similar in both occupations,[7] in contrast to tax avoidance effects that predominantly affect the self-employed. Hence, we tentatively predict a positive effect from income tax rates on self-employment rates.

In contrast to the effects of income tax, the effects of payroll taxes on the rate of self-employment have been less frequently studied. A high rate of payroll tax might induce employers to utilize self-employed contractors as a means of reducing the cost of labor, thus leading to a positive relationship between the payroll tax rate and the rate of self-employment. On the other hand, an increase in the rate of payroll tax may serve to reduce the incentive for self-employment among those who anticipate the need to hire other workers in order to run their business. Empirical evidence on this issue is limited. Using microdata on individuals in the United States, Moore (1983) finds a positive effect of payroll tax rates on the probability of self-employment. OECD (1992) reports a correlation coefficient of +0.7 between the rate of employers' social security contributions and the rate of self-employment in a cross-section of 19 OECD countries. Given the theoretical arguments discussed above, there are reasons to believe that the effects of payroll taxes may differ between the self-employed who employ others (owner-managers) and independent sole traders. Unfortunately, sufficient data that would enable us to test this proposition are not available.

The replacement rate is a potentially important variable as a high level of unemployment benefits might discourage unemployed workers from setting up in business for themselves. Moreover, as self-employed workers often do not enjoy the same benefit entitlements as those in waged employment, a high replacement rate could also discourage some workers from leaving paid-employment for self-employment for fear of losing their access to benefits. Staber and Bogenhold (1993) report a negative relationship between unemployment benefits and the self-employment rate in their analysis of OECD self-employment rates. We also predict a negative effect from replacement rates on self-employment rates.

Data on the level of per capita GDP (Y), the female labor force participation rate (F), the service sector share of GDP (V), the unemployment rate (U), the average income and payroll tax rates (A and P), and the replacement rate (R) were compiled from a variety of OECD publications. The Appendix provides detailed information on data definitions and sources. Means and standard deviations of the explanatory variables, along with those for the rate of nonagricultural self-employment, are displayed in Table 3. Unfortunately, as is apparent from Table 3, we were unable to obtain consistent time-series data on all of the explanatory variables for all of the countries in our sample. Consequently, in our econometric analysis we estimated a model using data over 1972–1993 for five countries: Finland, Japan, Sweden, the United Kingdom, and the United States. This contains examples from all three groups of countries illustrated in Figure 1.

The salient features of the explanatory variables may be briefly described. Table 3 shows quite wide variation in the mean values of per capita GDP and rates of unemployment across the countries in our sample. Rates of female labor force participation also show quite wide disparities, being particularly low in Ireland, Italy, and Spain, and relatively high in Sweden. There is somewhat less variation across countries in the service sector share of GDP, which ranges from 55% in Japan to just under 67% in the United States. The average rate of income tax stands out as particularly high in Sweden (54.3%) and Finland (42.7%). For the rate of payroll tax, the countries may be divided neatly into those with low rates of tax (the United States, Japan, and the United Kingdom) and those with much higher rates (Finland, France, Italy, and Sweden). Finally, the five countries for which we have consistent time series data on the replacement rate may again be divided into those with relatively low

[7] For supportive evidence from the United Kingdom, see Ajayi-obe and Parker (2005).

Table 3. Summary Statistics

	S	Y	U	F	V	A	P	R
Descriptive statistics								
United States	6.77	16.35	6.61	62.36	66.79	24.49	6.90	12.52
	(0.30)	(1.76)	(1.26)	(6.79)	(3.56)	(1.34)	(0.83)	(1.63)
Canada	6.56	14.94	8.73	60.23				
	(0.79)	(1.97)	(1.90)	(7.56)				
Japan	12.23	11.63	2.30	57.05	55.17	19.92	8.50	10.43
	(1.44)	(2.57)	(0.52)	(3.57)	(3.21)	(3.97)	(1.60)	(1.92)
Australia	11.17	13.34	6.88	55.70	62.81	17.14		
	(0.88)	(1.55)	(2.46)	(5.88)	(4.84)	(1.45)		
Finland	6.83	11.42	6.91	70.05	59.70	42.71	23.33	28.73
	(1.09)	(1.60)	(5.11)	(3.21)	(4.67)	(7.25)	(2.84)	(8.64)
France	9.34	12.30	7.98	55.02	64.10	37.32	29.63	
	(1.13)	(1.39)	(3.12)	(3.01)	(5.08)	(5.23)	(2.27)	
Ireland	9.91	7.79	11.98	38.45				
	(0.97)	(1.96)	(4.16)	(4.74)				
Italy	19.24	10.75	9.23	40.44	58.51	30.06	26.56	
	(1.31)	(1.71)	(2.31)	(4.03)	(5.82)	(4.61)	(2.82)	
Norway	6.37	13.20	3.17	64.86	62.73			
	(0.63)	(2.63)	(1.62)	(7.77)	(4.38)			
Spain	14.15	8.20	14.12	36.54	58.04			
	(0.63)	(1.21)	(7.29)	(4.93)	(4.17)			
Sweden	5.71	13.08	3.32	74.25	65.76	54.30	26.32	23.49
	(1.57)	(1.18)	(2.24)	(5.44)	(3.72)	(6.50)	(6.47)	(8.16)
United Kingdom	8.76	11.34	7.41	60.17	64.35	24.26	6.99	21.65
	(1.93)	(1.64)	(3.09)	(4.77)	(4.43)	(2.19)	(0.66)	(2.97)
Panel unit root statistics								
ψ_t	0.59	−1.31	−0.58	3.66	−1.10	−0.02	−1.39	−0.53
$\psi_{\overline{LM}}$	−1.05	−0.61	1.24	1.78	0.55	0.87	5.50	1.59

S = nonagricultural self-employment rate (%); Y = real per capita GDP (in thousands of 1985 U.S. dollars); U = unemployment rate (%); F = female labor-force participation rate (%); V = value added in services as a percentage of GDP; A = average rate of tax; P = payroll tax rate (%); R = replacement rate (OECD summary measure, %). See the Appendix for precise definitions of the variables and data sources. Descriptive statistics are sample means, with standard deviations in parentheses. Gaps indicate where data are unavailable on a consistent basis over the entire sample period. This is 1972–1996 for all variables except *R*, for which it is 1972–1993. The panel unit root statistics ψ_t and $\psi_{\overline{LM}}$ are described in the text. They are both distributed as standard normal variates under the null hypothesis of no panel unit root. They are calculated for the natural logarithm of each variable and are used for one-tail tests, with rejection of the null being in the left-hand tail.

rates (the United States and Japan) and those with somewhat more generous unemployment benefits (Finland, Sweden, and the United Kingdom).

3. Panel Unit Roots and Cointegration: Tests and Results

This section asks whether a long-run equilibrium relationship exists between self-employment rates and the proposed explanatory variables. This requires the use of panel data integration and cointegration tests. Because these are relatively new developments in the econometric literature, we briefly describe them below.

Henceforth, we use the following notation: let $i = 1, \ldots, N$ index the different countries in the panel, and let $t = 1, \ldots, T$ index time. Let s_{it} denote country i's (log) self-employment rate at time t. There are M explanatory variables (in logs) indexed by $j = 1, \ldots, M$, and denoted by x_{jit}.

Panel Unit Root Tests

In order to investigate the possibility of panel cointegration, it is first necessary to determine whether self-employment rates and the explanatory variables evolve as unit root processes.

Among the best-known panel unit root tests are those of Im, Pesaran, and Shin (1997; henceforth IPS), which are based on the well-known Dickey-Fuller procedure. For any variable $y_{jit} \in \{s_{it}, x_{1it}, \ldots, x_{Mit}\}$, the IPS tests involve estimating

$$\Delta y_{jit} = \xi_{ji} + \phi_{ji} y_{ji,t-1} + \gamma_{ji} t + \sum_{l=1}^{g_{ji}} \rho_{jil} \Delta y_{ji,t-\ell} + v_{jit} \quad t = 1, \ldots, T, \tag{1}$$

for each country i, where ξ_{ji} is a country-specific intercept, and g_{ji} are the number of lagged dependent variables required to rid the disturbances v_{jit} of serial correlation. The null hypothesis is that $\{y_j\}$ has a unit root: that is, $\phi_{ji} = 0$ for all i. The alternative is that $\{y_j\}$ is trend stationary for at least some of the countries; that is, $\phi_{ji} < 0$ for some i.[8]

IPS proposed two test statistics: the t-bar and LM-bar statistic. The t-bar statistic is constructed as follows: Denote the t-ratio of $\hat{\phi}_{ji}$ from Equation 1 by $t_{jiT}(g_{ji})$. Then the t-bar statistic is the average of these t-ratios across the countries:

$$\bar{t}_{jNT} = \frac{1}{N} \sum_{i=1}^{N} t_{jiT}(g_{ji}).$$

IPS proved that the following standardized t-bar statistic converges to a standard normal variate:

$$\Psi_{j,\bar{t}} = \frac{\sqrt{N}[\bar{t}_{jNT} - \frac{1}{N}\sum_{i=1}^{N} \mu(g_{ji}, T)]}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} \sigma^2(g_{ji}, T)}} \quad \Rightarrow N(0, 1),$$

where $\mu(g_{ji}, T)$ and $\sigma^2(g_{ji}, T)$ are constants tabulated in Table 2 of IPS (1997). Rejection of the null hypothesis of a panel unit root occurs in the left-hand tail of the standard normal distribution.

The LM-bar statistic is obtained as follows. Define

$$LM_{jiT}(g_{ji}) = \frac{T \Delta \mathbf{y}_{ji}' \mathbf{P}_{ji} \Delta \mathbf{y}_{ji}}{\Delta \mathbf{y}_{ji}' \mathbf{M}_{ji} \Delta \mathbf{y}_{ji}},$$

where $\mathbf{M}_{ji} = \mathbf{I} - \mathbf{Q}_{ji} (\mathbf{Q}_{ji}'\mathbf{Q}_{ji})^{-1} \mathbf{Q}_{ji}'$, $\mathbf{Q}_{ji} = (\mathbf{i}, \mathbf{t}, \Delta\mathbf{y}_{ji,-1}, \ldots, \Delta\mathbf{y}_{ji,-g_{ji}})$, $\mathbf{i}$ is a vector of ones, $\mathbf{t}$ is the time trend, $\Delta\mathbf{y}_{ji} = (\Delta y_{ji1}, \Delta y_{ji2}, \ldots, \Delta y_{jiT})'$, $\mathbf{P}_{ji} = \mathbf{M}_{ji}\mathbf{y}_{ji,-1} (\mathbf{y}_{ji,-1}'\mathbf{M}_{ji}\mathbf{y}_{ji,-1})^{-1} \mathbf{y}_{ji,-1}'\mathbf{M}_{ji}$, and $\mathbf{y}_{ji,-1} = (y_{ji0}, y_{ji1}, \ldots, y_{ji,T-1})'$ Then, analogous to the t-bar statistic, the LM-bar statistic is $\overline{LM}_{jNT} = (1/N) \sum_{i=1}^{N} LM_{jiT}(g_{ji})$, and the following standardized LM-bar statistic converges to a standard normal variate:

$$\Psi_{j,\overline{LM}} = \frac{\sqrt{N}\left[\overline{LM}_{jNT} - \frac{1}{N}\sum_{i=1}^{N} \tilde{\mu}(g_{ji,}T)\right]}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} \tilde{\sigma}^2(g_{ji,}T)}} \quad \Rightarrow N(0, 1),$$

where $\tilde{\mu}(g_{ji}, T)$ and $\tilde{\sigma}^2(g_{ji}, T)$ are constants tabulated in Table 1 of IPS. As with the standardized t-bar statistic, rejection of the null is in the left-hand tail of the standard normal distribution.

[8] As IPS point out, this is a more general alternative hypothesis than constraining all ϕ_{ji}, for $i = 1, \ldots, N$, to be less than zero, as studied by Levin and Lin (1993). It is noteworthy in this respect that Karlsson and Löthgren (2000) show that the IPS tests are more powerful than Levin and Lin's. See the special November 1999 issue of the *Oxford Bulletin of Economics & Statistics* for a comparison of panel unit root tests.

The bottom panel of Table 3 presents the standardized t-bar and LM-bar panel unit root statistics for each variable described in Section 2. There are only 25 time series observations per country, so the power of these tests is limited, despite gaining power relative to single country tests by pooling data across the panel.[9] It is notable that the null hypothesis of a panel unit root cannot be rejected for any variable. This implies that previous studies based on least squares estimation may be vulnerable to the spurious regression problem and motivates the use of panel data cointegration methods. We proceed cautiously on this basis, in view of the limited sample sizes available.

Panel Cointegration Tests

Consider the following regression:

$$s_{it} = \alpha_i + \delta_i t + \beta_{1i} x_{1it} + \beta_{2i} x_{2it} + \ldots + \beta_{Mi} x_{Mit} + e_{it} \quad \text{for } t = 1, \ldots, T; \quad i = 1, \ldots, N. \qquad (2)$$

The α_i parameters are country-specific intercepts, or fixed effects, and the $\delta_i t$ terms allow for country-specific time trends. It should be noted that the slope coefficients β_{1i}, β_{2i}, ... β_{Mi} are permitted to vary across individual countries in the panel. All variables in Equation 2 are assumed to be unit root processes.

Pedroni (1999a) has proposed two types of panel cointegration statistics designed to test the null hypothesis of no cointegration between the variables in Equation 2 against the alternative hypothesis of cointegration. Let γ_i denote the autoregressive coefficient of the estimated residuals for country i, $i = 1, \ldots, N$. The null hypothesis for both types of statistic is the same; that is, H_0: $\gamma_i = 1\ \forall i$. The alternative hypothesis for the first type of statistic is H_1: $\gamma_i = \gamma < 1\ \forall i$. The alternative hypothesis for the second type of statistic is less restrictive, being H_1: $\gamma_i < 1\ \forall i$. Because the latter allows an additional source of heterogeneity across individual countries in the panel, this type is the most general, for which results are reported below.[10]

We focus on two panel cointegration statistics of the second type proposed by Pedroni (1999b). The first is analogous to the Phillips and Perron (1988) t-statistic, and the second is analogous to the augmented Dickey–Fuller (ADF) t-statistic. Both statistics incorporate corrections for heteroscedasticity between countries and autocorrelation within countries, allowing the long-run covariance matrix to vary across the panel:

$$N^{-1/2}\tilde{Z}_{t_{N,T}} := N^{-1/2} \sum_{i=1}^{N} \left(\hat{\sigma}_i^2 \sum_{t=1}^{T} \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{t=1}^{T} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i)$$

$$N^{-1/2}\overline{Z}^*_{t_{N,T}} := N^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{s}_i^{*2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{t=1}^{T} \hat{e}^*_{it-1} \Delta \hat{e}^*_{it},$$

where $\hat{e}_{it}$ are the residuals estimated from Equation 2, $\hat{\lambda}_i$ is a set of estimated nuisance parameters derived from the long-run covariance matrix, and $\hat{\sigma}_i^2$ is the corrected variance of the autoregression residuals. The $\hat{e}^*_{it}$ are autocorrelation-corrected residuals, and $\hat{s}_i{*}^2$ are their variances (see Pedroni 1999b for details).

Pedroni (1999b) shows that following an appropriate standardization, these statistics are

[9] An advantage of the IPS test is that it allows lag lengths to vary across countries within each test. We do not report the lag lengths for each country and each variable for brevity. We also suppress the results obtained using an alternative panel unit root test suggested by Maddala and Wu (1999), for which similar results were obtained. All of these results are available from the authors on request.

[10] An alternative panel cointegration test has been proposed by Larsson, Lyhagen, and Löthgren (1998). This is based on an extension to the panel context of Johansen's (1988) multivariate likelihood-based approach. However, as noted by Banerjee (1999), the status of this test is unclear since the distribution of their cointegration statistic is only asserted, rather than proven, to have the standard normal distribution.

asymptotically distributed as standard normal variates. We denote the standardized statistics by PP-t and ADF-t, respectively. Both diverge to negative infinity under the alternative hypothesis of panel cointegration. For the five countries for which data on every explanatory variable described in Section 2 was available, the PP-t statistic took the value -1.33 ($p = 0.09$), whereas the ADF-t statistic took the value 6.63 ($p < 0.01$). The latter indicates strong evidence of cointegration, unlike the former. But in view of Monte Carlo studies that have shown the superiority of ADF-based unit root tests over alternatives (including PP; see Banerjee et al. 1993), we interpret this as providing support for the notion of cointegration between self-employment rates and the hypothesized variables.

Results of Estimating the Cointegration Vectors

We estimate here the cointegrating relationship between self-employment rates and the explanatory variables. We do this by pooling the long-run information in the panel while allowing the short-run dynamics and fixed effects to be heterogeneous among the different members of the panel. We seek the pooled long-run estimates of the beta coefficients linking self-employment rates to the explanatory variables.

It is now well known that OLS is a biased and inconsistent estimator when applied to cointegrated panels (see footnote 2 for references). To illustrate how serious this bias can be, column 1 of Table 4 presents OLS fixed-effects estimates for each of the five models.[11] The R^2 was 0.94, which exceeded the Durbin-Watson autocorrelation statistic of 0.81. This is a well-known indicator of spurious regression. Indeed, as shown by Pedroni (1999b), not only are the OLS coefficients biased when variables are nonstationary, so are their distributions, ruling out valid inferences based on their computed t-statistics.

To tackle this problem, Pedroni (1999b) proposed a fully modified OLS estimator (FMOLS) that provides consistent estimates of the beta coefficients, together with "t-ratios" that are asymptotically distributed as standard normal variates.[12] As with the panel cointegration test statistics, a correction for cross-panel heteroscedasticity and autocorrelation is applied to the estimator.

Column 2 of Table 4 presents estimates of the cointegration vectors and t-ratios for the model. The results are broadly consistent with our priors. Higher average tax rates and lower female participation rates and replacement rates are significantly associated with higher self-employment rates in the panel, as predicted. The effects from value added in services and payroll taxes are positive and negative, respectively, but neither influence is statistically significant. Neither of the other two macroeconomic variables used in previous work (the unemployment rate and per capita GDP) is statistically significant.

Given the relatively restricted sample of observations available for estimation of our most complete model specification, we should be careful of reading too much into these results.[13] Nonetheless, the results contain several implications for current and future work on the determinants

[11] We are grateful to an anonymous referee and Andrew Oswald for suggesting the inclusion of year dummies, as well as country-specific fixed effects dummies, in these regressions. Results were obtained using Version 4.3 of the Time Series Processor (TSP).

[12] Year dummies are redundant in this model, and therefore were excluded. Pedroni (1999b) proposed two types of t-statistic to test the null hypothesis of a single parameter restriction. The most general type is termed "group mean fully modified statistics": these are reported in columns 2, 3, and 4 of Table 4. Pedroni shows that significance tests based on these statistics are powerful and well-sized for $T \geq N$, which is the case here.

[13] There is no reason to suspect that the particular selection of countries available for the estimation of model 5 has any significant bearing on the results. As the discussion in section 2 demonstrates, these countries represent a balanced sample in terms of the values both of the dependent variable and the explanatory variables in the model.

Table 4. Unmodified OLS and Fully Modified OLS (FMOLS) Parameter Estimates

Coefficients	Unmodified OLS (1)	FMOLS I (2)	FMOLS II (3)	FMOLS III (4)
Real GDP per capita	−0.97***	0.78	0.76**	0.54*
	(3.73)	(1.39)	(2.02)	(1.75)
Unemployment rate	−0.09	0.00	−0.24*	−1.49
	(1.56)	(0.29)	(1.89)	(1.39)
Female participation rate	−2.09***	−1.41**	−1.42***	−1.26***
	(3.54)	(2.17)	(2.74)	(2.85)
Value added (services)	0.36	0.47	0.41	1.15
	(0.64)	(0.92)	(1.13)	(0.36)
Average rate of tax	−0.86***	0.60***	0.48***	0.95***
	(5.04)	(3.19)	(3.64)	(5.33)
Payroll tax	0.16	−0.28	−0.23	−0.56*
	(0.15)	(0.26)	(0.46)	(1.95)
Replacement rate	0.05	−0.23***	−0.24***	−0.73
	(0.82)	(3.82)	(4.97)	(1.39)
Unemployment squared			0.11*	
			(1.68)	
Unemployment–replacement rate interaction				0.43
				(1.55)
Sample size, n	110	110	110	110

Dependent variable is S, the pooled log self-employment rate. Absolute t-statistics are in parentheses. FMOLS I is the basic specification; FMOLS II augments it with a squared unemployment rate term, and FMOLS III augments it with an unemployment-replacement rate interaction. All three of these models are estimated by fully modified ordinary least squares (FMOLS). See the text for details.

* Denotes statistical significance with a type I error of 10%, ** of 5%, and *** of 1%.

of self-employment both within and across countries. First, notice that by comparing the results in columns 1 and 2 of Table 4, conventional panel data OLS estimation would have failed to obtain these results. Second, taxes and benefits (and also the female labor-force participation rate), rather than the macroeconomic variables of per capita GDP, unemployment, or aggregate industrial structure, appear to be the salient influences on self-employment rates in the OECD countries under study. This is of interest for two reasons. First, previous empirical work has emphasized macroeconomic factors, rather than tax and benefit variables. Our findings suggest that it might be necessary to reconsider the econometric specifications used in future research. Second, because tax and benefit variables are under direct government control, governments may have considerable influence on the extent of self-employment within their economies. Although the promotion of self-employment is unlikely to be a primary objective of government tax and benefit policies, governments should be aware of the implications for self-employment of their policy decisions in this area.

The results can also shed light on specific cross-country differences. For example, Sweden and Finland are countries with some of the lowest self-employment rates. Yet these countries are also among the highest for female participation and benefit replacement rates. The opposite is the case for Japan and the United Kingdom, which have lower female participation and benefit replacement rates. The United States is in an intermediate position, with average tax rates and labor force participation rates. That the differences between average self-employment rates in Japan and the United Kingdom on the one hand, and Finland and Sweden on the other, are not greater seems to be chiefly attributable to offsetting effects from average tax rates. These are predicted to bolster self-employment rates most in Finland and Sweden and least in Japan, the United States, and the United Kingdom. Regarding trends, the greatest increases in self-employment rates were observed in the United Kingdom and Sweden. Both countries witnessed especially pronounced reductions in replacement rates over the sample period. In contrast, both benefit replacement rates and female labor-force participation rates increased in Japan. This can partly explain the downward self-employment trend in Japan. In contrast, changes in explanatory variables appeared to offset each other in the United States, where little self-employment trend was observed.

Finally, we explored whether we can explain the mixed empirical results in the literature relating to the effect of unemployment on self-employment rates. To shed light on this issue, we included a quadratic unemployment rate variable in the specification. The reason is that at high rates of unemployment, there are numerous workers available to start new businesses. The opposite is the case in economies with lower unemployment rates. However, economies with very low unemployment rates might enjoy this position partly because their citizens have (unobserved) probusiness attitudes that translate into a strong predisposition for both work and entrepreneurship. Then one might expect to see a *U*-shaped relationship between self-employment and unemployment rates. This is precisely what we find in column 3 of Table 4. Japan has the lowest and the United Kingdom has the highest unemployment rates in our sample; they also have the highest self-employment rates. The other three countries all have lower unemployment *and* lower self-employment rates.

We also tried interactions between unemployment rates and some other variables. In principle, we might expect the push effect of unemployment on self-employment to be stronger in countries with relatively low benefit replacement rates. There would be greater incentives in such countries to exit unemployment for self-employment for any given level of aggregate demand. We tested this possibility by including as an additional explanatory variable an interaction between replacement and unemployment rates. The results appear in column 4 of Table 4. Quite apart from the statistical insignificance of the interaction term, they do not support the notion that different replacement rates in different countries explain different effects of unemployment on self-employment. The coefficients take the opposite signs (−1.49 and +0.43) to those expected under the hypothesis. Finally, we tried including an interaction of unemployment with GDP per capita, the idea being that unemployment might have a greater positive effect on self-employment in countries with high per capita demand (GDP). However, the results were similarly unsupportive, and for the sake of brevity are not reported here.[14]

4. Conclusion

This article has built on previous studies of the determinants of self-employment by using a wider range of variables and by applying recently developed panel cointegration techniques to a panel of OECD countries. Special emphasis was placed on the possible role of government policy instruments in affecting the self-employment rate. Unlike single-country studies, a panel approach allows comparisons between countries at similar points in time as well as within countries over time. Panel data also increases the power of cointegration tests. We first documented the considerable variety in self-employment rates and trends within major OECD economies between 1972 and 1996. We then presented evidence that a set of explanatory variables cointegrates with self-employment

[14] Another possibility is that government regulations and liquidity constraints (e.g., Blanchflower and Oswald 1998) also discourage unemployed people from switching into self-employment. The unavailability of such variables in our sampling frame means that our results should probably be treated with caution. But this should not deter future researchers from seeking ways of refining the analysis, possibly by using more detailed national data sets.

rates in a panel of five countries before showing that several macroeconomic variables proposed in previous studies do not appear to explain robustly the evolution of international self-employment rates. Rather, tax-benefit variables and the female labor-force participation rate are found to possess most of the explanatory power. Of particular interest is the finding that self-employment rates are positively and significantly related to average income tax rates, and negatively and significantly related to the benefit replacement rate. These findings suggest a stronger influence of government policy decisions in the determination of cross-national variations in self-employment rates than has typically been recognized in the literature to date. Although subject to certain limitations acknowledged in the text, our findings appear robust to different model specifications. We also showed that conventional panel data OLS estimation would have failed to identify these results.

Appendix

S—Rate of nonagricultural self-employment. Nonagricultural self-employed as a percentage of total civilian nonagricultural employment (employees plus self-employed) plus unemployment. Source: *OECD Labor Force Statistics*.

Y—Per capita real GDP. GDP per capita in constant (1985) dollars, expressed in international prices. Sources: Penn World Tables (1972–1992), updated to 1996 using information from OECD country surveys and *National Accounts Statistics*.

U—Rate of unemployment. Unemployed as a percentage of the labor force. Source: *OECD Labor Force Statistics*.

F—Female labor-force participation rate. Total female labor force as a percentage of the female population aged 15–64. Source: *OECD Labor Force Statistics*.

V—Value added in services as a percentage of GDP. Source: *OECD Main Economic Indicators, Historical Statistics* and *OECD National Accounts Statistics*.

A—Average rate of income tax. Income tax payments plus employees' social security contributions as a percentage of wages and salaries plus the operating surplus of unincorporated enterprises plus property income. Sources: Tables of the income and outlay account of households and unincorporated enterprises in *OECD National Accounts Statistics* and *UN National Accounts Statistics*.

P—Payroll tax rate. Employers' contributions to social security as a percentage of wages and salaries. Sources: Income and outlay account of households and unincorporated enterprises, *OECD National Accounts Statistics* and *UN National Accounts Statistics*.

R—Replacement rate. Overall average of gross replacement rates for three types of families (single person, married person with dependent spouse, married person with spouse in work) and two earnings levels (average earnings and 66.7% average earnings). Source: *OECD Database of Benefit Entitlements and Gross Replacement Rates*. Data kindly supplied by John Evans of OECD. Note that the OECD data are compiled on a biennial basis; figures for missing years have been interpolated.

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